

Theories of cartel stability and the Joint Executive Committee

Glenn Ellison*

This article reexamines the experience of the Joint Executive Committee, an 1880s railroad cartel, to assess the applicability of the Green and Porter (1984) and Rotemberg and Saloner (1986) theories of price wars. After discussing necessary modifications to the theories, I estimate a number of dynamic models to explore the causes of price wars, the cyclical nature of pricing, and the possibility that secret price cuts may have been given. The estimates provide some support for the predictions of the first theory.

1. Introduction

■ The “Joint Executive Committee” (JEC) was a railroad cartel organized in 1879 to set prices for transport between Chicago and the East Coast. Over the next seven years the cartel was only partially successful, with rates stable at \$7 or \$8 per ton in some years and occasional price wars pushing rates as low as \$2.50 per ton in others. One thing the cartel did do very well was to keep detailed records, and as a result the JEC has proved to be a valuable setting in which to study the workings of a cartel (MacAvoy, 1965; Ulen, 1979; Porter, 1983b). In light of the previous literature, I will not review the history of the cartel here and recommend the accounts of MacAvoy and Porter to anyone unfamiliar with its situation.

Much of the more recent attention paid to the JEC is attributable to interest in game-theoretic models of collusion. Green and Porter’s (1984) analysis of collusion with imperfect observation demonstrated that price wars need not be indicative of a failure to collude; they are, in fact, a necessary part of optimal collusive arrangements. The debate was further energized by the work of Rotemberg and Saloner (1986), whose alternate description of price wars is often regarded as at odds with that of Green and Porter. Proponents of each theory have cited the JEC’s experience as lending empirical support. (Porter, 1983b; Lee and Porter, 1984; Rotemberg and Saloner, 1986). Although the JEC has received a great deal of attention, I shall argue that the existing literature allows only a fairly incomplete assessment of the applicability of the Green and Porter and Rotemberg and Saloner theories. In an attempt to provide a more complete picture I shall reexamine the JEC, discussing at length the implications of the theories given the situation of the JEC and performing a number of empirical tests.

* Harvard University.

I would like to thank Sara Fisher Ellison, Drew Fudenberg, Wally Mullin, Danny Quah, Julio Rotemberg, and Jean Tirole for their comments and suggestions. Special thanks are due to Rob Porter both for his helpful suggestions and for providing me with the data used in this article.

The first theory considered here, that of Green and Porter, has been the primary focus of most previous studies. These studies have provided convincing evidence that price wars did occur in the JEC. The core content of the theory, however, is not just that price wars should occur in equilibrium, but that price wars should occur because firms follow particular strategies to sustain collusion, strategies in which price wars are triggered by random demand shocks. To evaluate this prediction within the context of the JEC, it is important to note that a variety of trigger strategies could have been used to sustain a nearly optimal outcome and that these strategies may be somewhat different from those in the text of Green and Porter (1984). For example, firms might react to unusual market share patterns or begin a price war whenever demand is unusually high (as opposed to low). The empirical analysis focuses on the transition probabilities in a switching regressions model (incorporating elements from the models of Porter (1983b), Cosslett and Lee (1985), and Hajivassiliou (1989)) to see whether any such trigger strategies were used. Some evidence for this is found.

The second theory considered, that of Rotemberg and Saloner (1986), is commonly associated with the statement that price wars are more likely to occur during booms, and therefore viewed as somehow in opposition to the Green and Porter theory. The actual Rotemberg and Saloner model, however, is really about countercyclical pricing—firms have perfect information and adjust prices smoothly in response to demand conditions. In contrast, the JEC was a cartel with imperfect information because secret price cuts were possible. When adapted to such a situation, the Rotemberg and Saloner analysis might predict either countercyclical pricing or price wars during booms. Interestingly, these predictions may be applied not only to booms in the traditional sense, but also to the seasonal pattern in the JEC's demand. Here, however, the empirical evidence provides little support for the theory.

Finally, the article considers a simple but heretofore unexplored prediction common to each of these theories, that firms should not cheat on their agreements. The discussion of secret price cutting requires the estimation of a set of more complicated hidden regime models, and stresses the problems of interpretation inherent in such models. While some cheating apparently occurred, it appears to have been fairly rare. This may be taken as further supporting the Green and Porter theory to the extent that we would expect rational firms to have cheated if the proper dynamic incentives to cooperate were not maintained.

2. The Green and Porter theory

■ **Theory and previous results.** Green and Porter (1984) first formalized the observation that collusion is possible with imperfect information in a model of repeated Cournot competition between identical firms. They assume production levels are unobservable, and that there is noise in demand. Firms do observe the market price, and therefore receive an imperfect signal of their competitors' behavior. What Green and Porter show is that a degree of collusion can be sustained via trigger strategies that involve switches between collusive periods and price wars. In a collusive period, the firms produce at a point between the Cournot and monopoly levels. In a price war, the firms produce at the Cournot level. The optimal degree of collusion is sustained by trigger strategies in which firms produce at the collusive level until the market price falls below a specified level, at which point a price war begins. Abreu, Pearce, and Stachetti (1986) extend the analysis by allowing for more general strategies and show that an equilibrium involving two states and trigger strategies is indeed optimal.

In two important respects the JEC differs from the classic Green and Porter formulation. First, it is certainly more reasonable to think of the railroads not as setting quantities, but rather as setting prices with reputations for quality of service and differing route structures creating differentiated demands. Second, what is and is not observable is also

different. Given the possibility of secret price cuts and with no single market price, price is not observable. On the other hand, the JEC did collect quantity data that presumably were reliable.¹

In developing predictions applicable to the JEC, it is probably best not to push too far the analysis of optimal equilibria. With asymmetric firms and observable market shares, the optimal equilibrium may well resemble those of Fudenberg, Levine, and Maskin (1993), where full collusion is supported by equilibria that specifically punish deviators. If we tried to apply such a theory, however, we would be immediately faced both with the reality that no asymmetric punishments were observed and with limitations in the amount of data available to identify complex strategies. Even restricting ourselves to symmetric trigger strategies, what is optimal may also be hard to determine without knowing far more details than are available.

If we ask instead what simple, nearly efficient equilibria are possible, we can find a number of plausible triggers upon which an equilibrium might be based. The best signal of cheating with observable market shares would likely be some pattern of a high market share for one firm with lower market shares for all the rest. A second possible signal is *high* aggregate demand, as when one or more firms lower prices there is not only a transfer of market shares but also an expansion of demand. (This contrasts with the low aggregate demand that causes price wars in the literal Green and Porter model.) Finally, if we wished to find an equilibrium robust to some type of coalition-proofness notion, it might make sense for firms to react to an unusually low market share of one firm, because a firm will experience unusually low sales when one or several other firms cheat.

In trying to assess whether the organization of the JEC resembles any of these equilibria, we may test several predictions. First, price wars should occur. Second, price wars should occur precisely when random demand shocks resemble a signal like those noted above. Finally, firms should not cheat and offer secret price cuts. I defer any further discussion of this last issue until the final section.

The previous empirical work on the JEC has largely concerned the first of these predictions. Porter (1983b) first showed that the data are indicative of multiple supply regimes with prices were reduced by about 40% during price wars. Cosslett and Lee (1985) noted further that the price wars do indeed appear to be continuous rather than a randomly scattered set of weeks. Lee and Porter (1984), Berry and Briggs (1988), and Hajivassiliou (1989) provide additional evidence on the existence of price wars.

While central to the theory, the second prediction has been addressed only in Porter (1985). Perhaps further work has been discouraged by Bresnahan's (1989) observation that with only six to eleven price wars to work with, we have much less data on such dynamic issues. Porter uses a two-step approach in which a collusion indicator is first estimated by maximum likelihood and then regressed on a measure of deviations of market shares from the assigned quotas over the sample of collusive periods. While Porter does not find significant evidence of such a trigger, we cannot conclude that no trigger exists. There are many reasonable triggers for which Porter does not look. In addition, the two-step approach is perhaps susceptible to a loss of power because the onset of price wars may be misclassified in the first stage. Interestingly, Porter does find that large aggregate demand (he does not test residuals) does appear to cause price wars. In his framework, however, it is not possible to tell if this is a trigger strategy or related to Rotemberg and Saloner's predictions.

□ **Model and data.** The models of this article are all based on that of Porter (1983b). The primary difference is that I impose a Markov structure related to those of Cosslett

¹ Some support for this assumption is provided by the cartel's consideration of plans in which the shipment data were to be used to determine side payments by firms exceeding allotments.

and Lee (1985) and Hajivassiliou (1989) on the transitions between the cooperative and price war regimes. I test whether collusion was maintained by trigger price strategies by focusing on the determinants of the probability with which a price war begins.

The structural model described below follows Porter's specification very closely, so it will only be sketched here. The reader may refer to Porter (1983b) for motivation and interpretation of the parameters. The demand for grain shipments is assumed to be given by

$$\log Q_t = \alpha_0 + \alpha_1 \log P_t + \alpha_2 LAKES_t + \alpha_{3-14} SEASXX_t + U_{1t}, \quad (1)$$

where $SEASXX_t$ and $LAKES_t$ are intended to capture seasonal variation in demand and the effect of the opening and closing of the Great Lakes to navigation. In a departure from Porter's model, I will usually assume that U_{1t} follows an AR(1) process,

$$U_{1t} = \rho U_{1t-1} + V_{1t} \quad |\rho| < 1. \quad (2)$$

The supply curve, also as in Porter, is assumed to have the form

$$\log P_t = \beta_0 + \beta_1 \log Q_t + \beta_2 I_t + \beta_{3-6} DMx_t + U_{2t}, \quad (3)$$

where I_t is an indicator of collusion, and the DMx are four dummy variables for changes in the cartel composition. The errors V_{1t} and U_{2t} are assumed to have a multivariate normal distribution. The possibility of secret price cuts makes P_t unobservable, but we do have available the official prices firms were supposed to have charged. Although using this series could introduce an errors-in-variables problem, the optimal cartel theory predicts that firms should in fact never deviate from the official prices, so we can hope that these errors are small.

As mentioned above, I use a Markov structure on the transitions to detect causes of price wars. Specifically, I assume that the indicator I_t of collusion evolves according to the logit model

$$\text{Prob}\{I_{t+1} = 1 \mid I_t, Z_t\} = \frac{e^{\gamma W_t}}{(1 + e^{\gamma W_t})}, \quad (4)$$

where Z_t is the set of predetermined variables at time t . Taking W_t to be a constant independent of I_t gives a standard model with independent regime switches, as in Porter (1983b). Allowing W_t to contain I_t adds a Markov structure to the price wars, as in Cosslett and Lee (1985), so that a noncooperative state today can be likely to lead to another noncooperative state next week. I shall also include several other variables in W_t to test the predictions of the Green and Porter model. Note that the specification differs from that of Hajivassiliou (who does not test for Green and Porter-style triggers) in that variables are modelled as affecting *transition* probabilities, i.e., the probability of starting a price war, not the probability of being in a given state.

Estimates of the parameters are computed by maximizing the joint likelihood of the structural equations and of the path of transitions. Hopefully, the simultaneous maximization will allow the model to date more accurately the start of price wars and provide more power in detecting triggers than would a two-step approach. The computation is similar to that of Cosslett and Lee (1985). I use numerical derivatives to compute robust standard errors for all estimators.

The data for the article are those compiled by Ulen (1979) and later used by Porter (1983b). The data contain weekly observations on quantity, price, and other variables over a 328-week period from 1880 to 1886. Table 1 gives summary statistics.

QUANTITY gives the total eastbound grain shipments of the railroads in the JEC in tons per week. *PRICE* is given in dollars per 100-pounds of grain. The *LAKES* variable is a dummy variable set to one when the Great Lakes were navigable and steamers could

TABLE 1 Summary Statistics

Variable	Mean	Standard Deviation	Minimum	Maximum
<i>QUANTITY</i>	25384	11632	4810	76407
<i>PRICE</i>	.2465	.0665	.125	.400
<i>LAKES</i>	.5732	.4954	0	1
<i>DM1</i>	.4238	.4949	0	1
<i>DM2</i>	.0457	.2092	0	1
<i>DM3</i>	.4329	.4962	0	1
<i>DM4</i>	.0152	.1227	0	1
<i>BIGSHARE1</i>	1.0908	.5690	.0397	2.9711
<i>BIGSHARE2</i>	1.2409	.7102	.1849	4.2345
<i>BIGSHAREQ</i>	1.1744	.5162	.1578	2.9728
<i>SMALLSHARE</i>	1.2303	.7373	.1161	5.7565

compete with the railroads. The four structural dummies *DM1–DM4* mark significant changes in the membership of the cartel. Seasonal dummy variables labelled *SEAS1–SEAS12* are used for the first twelve four-week periods of each year. All of these variables are exactly as in Porter (1983b).

In addition, I created four more complicated variables to test the Green and Porter model. The variables are admittedly ad hoc, and they represent a rough attempt to capture a few of the possible workings of a cartel. The variable *BIGSHARE1* is intended to be a plausible trigger in a cartel that switches to a price war when one firm obtains a suspiciously high market share. The measure is based on deviations in $\log q_{it}$ rather than q_{it} so that differing punishment probabilities will not disturb the proportionality between gains from a given increase in $\log q_{it}$ related to a price cut and losses from a subsequent price war. Recognizing that the measure is arbitrary, it seemed prudent to construct two variants on the theme, *BIGSHARE2* and *BIGSHAREQ*. Each of the variables is the largest amount by which the strength s_{it} of any firm's demand exceeds its predicted value \bar{s}_{it} (after normalizing by the sample standard deviation), i.e., each is of the form $\max_i(s_{it} - \bar{s}_{it})/\sigma_i$. In the case of *BIGSHARE1* and *BIGSHAREQ*, $s_{it} = \log q_{it} - \frac{1}{n} \sum_j \log q_{jt}$ is used to represent the strength of firm i 's demand. In the case of *BIGSHARE2*, the ordinary market share is used. The predicted values \bar{s}_{it} used in constructing *BIGSHARE1* and *BIGSHARE2* are taken to be the average of the same measure for the previous twelve weeks. The predicted value for *BIGSHAREQ* is computed from the assigned quota. While the use of quotas as expectations initially seems appealing, in practice the quotas were often set unrealistically, hence deviations will not accurately reflect deviations from adherence to a uniform price.² Adjustments to the predictions had to be made at the start of the series and as the cartel structure shifted. A fourth variable, *SMALLSHARE1*, is intended to be a plausible trigger for a cartel in which price wars are triggered by an unusually small market share for one firm. Using the same measures as in *BIGSHARE1*, it is the largest amount by which the strength of any firm's demand falls short of its predicted value.

□ **Results.** In this subsection I present first a "standard" model much like those that previous studies have used to identify price wars. Subsequently, I explore the causes of price wars. The results of the "standard" model are given in the left half of Table 2. The model differs from that of Porter (1983b) in two ways. First, it allows for serial correlation in the demand equation. The effect of this change can be seen by comparing the two halves of the table. Note that the coefficient U_{it-1} is highly significant and several other coef-

² For example, MacAvoy (1965) reports that the Chicago and Grand Trunk Railway was admitted with a 10% quota while receiving only a 2–7% share at the cartel price.

TABLE 2 The "Standard" Model

Variable	"Standard" Model		No Serial Correlation	
	Estimate	Standard Error	Estimate	Standard Error
Demand				
<i>CONSTANT</i>	7.677	1.882	9.019	.361
<i>log P</i>	-1.802	1.287	-.843	.193
<i>LAKES</i>	-.009	.112	-.460	.348
<i>SEAS1</i>	-.103	.086	-.117	.157
<i>SEAS2</i>	.146	.145	.167	.180
<i>SEAS3</i>	.147	.138	.149	.166
<i>SEAS4</i>	-.011	.157	.145	.242
<i>SEAS5</i>	-.315	.165	.062	.164
<i>SEAS6</i>	-.550	.179	.077	.170
<i>SEAS7</i>	-.446	.198	.081	.176
<i>SEAS8</i>	-.504	.194	-1.116	.374
<i>SEAS9</i>	-.395	.165	.048	.185
<i>SEAS10</i>	-.545	.164	.102	.191
<i>SEAS11</i>	-.521	.180	.085	.304
<i>SEAS12</i>	-.397	.173	.183	.241
Supply				
<i>CONSTANT</i>	-4.764	1.863	-5.649	9.461
<i>log Q</i>	.306	.178	.398	.928
<i>DM1</i>	-.154	.075	-.211	.124
<i>DM2</i>	-.246	.064	-.283	.160
<i>DM3</i>	-.317	.076	-.373	.242
<i>DM4</i>	-.198	.119	-.419	.422
<i>I_t</i>	.637	.104	.660	.406
Regimes				
<i>CONST. (I_t = 1)</i>	3.675	.474	3.661	.513
<i>CONST. (I_t = 0)</i>	-2.641	.404	-2.620	.476
Other				
σ_1	.290	.061	.396	.029
σ_{12}	-.007	.004	-.045	.142
σ_2	.160	.045	.191	.313
<i>U_{t-1}</i>	.832	.085		
Log-likelihood	181.0		37.2	

ficients change. Most importantly, the estimated price elasticity increases from .84 to 1.80, indicating that the JEC did indeed face an elastic demand curve. As a consequence, Porter's approximation θ to the degree of collusion increases from .40 to .85, indicating that the JEC was much more aggressive in its pricing than has been previously reported. (A value of one would indicate full monopoly pricing.) The second change is the incorporation of a Markov structure for the transition probabilities. Consistent with a number of previous studies, the regimes are not independently determined with collusion having probability .975 ($e^{3.67}/(1 + e^{3.67})$) after a collusive state and .067 after a price war state. On the whole, the model reinforces the view of the JEC as setting prices collusively between occasional price wars, suggesting a greater degree of collusion than has been claimed in the past.

I now turn to the second prediction described above, trying to infer whether price wars resulted from the use of Green and Porter-style trigger strategies by examining the causes of price wars. The results of six models are summarized in Table 3. Each model contains all of the variables of the "standard" model plus an additional variable (or two)

TABLE 3 Causes of Price Wars

Parameter	Model					
	1	2	3	4	5	6
Regimes: $\text{Prob}\{I_{t+1} = 1 \mid I_t = 1, Z_t\} = \frac{e^{\gamma W_t}}{(1 + e^{\gamma W_t})}$						
<i>CONSTANT</i>	4.63 (.77)	4.36 (.77)	3.95 (1.30)	2.96 (.66)	4.43 (.81)	4.35 (.90)
<i>BIGSHARE1</i>	-.77 (.49)					
<i>BIGSHARE2</i>		-.46 (.39)				
<i>BIGSHAREQ</i>			-.21 (1.06)			
<i>SMALLSHARE</i>				.66 (.89)		
V_{1t}					-4.15 (2.64)	-5.00 (3.15)
U_{1t}						1.17 (1.15)

* Note: Estimated standard errors in parentheses.

that affects the probability of a transition from a collusive to a price war state. For each variable, the theory predicts a larger value to be more likely to lead to a breakdown of collusion. The coefficient estimates are then expected to be negative. Because the “standard” model indicates that there are only six price wars, we should not expect high significance levels. (There are no substantial changes in the structural parameters of the models. Hence, only the logit parameters for the transition to a price war are reported.)

The first model in the table focuses on what should be the best signal of cheating, that of one firm obtaining an unusually high market share. The coefficient estimate on the *BIGSHARE1* variable is negative as predicted, and is significant at the 6% level in a one-tailed test. The next two columns of the table give estimates from models with the two other variants on this variable, *BIGSHARE2* and *BIGSHAREQ*. The estimates are again negative, but they are smaller and not significant. The fourth model includes the *SMALLSHARE* variable to test whether price wars are triggered by an unusually low demand for one firm. The coefficient estimate has the wrong sign and is not significant. The fifth model tests whether large residuals V_{1t} in aggregate demand cause price wars. The coefficient is large and negative, as the theory predicts, and is again significant at the 6% level. The fact that two essentially orthogonal triggers appear to increase the probability of a price war is also suggestive of the possibility that each is imperfectly correlated with some true trigger that has a more powerful effect.

The sixth model in Table 3 takes advantage of the incorporation of serial correlation to sharpen the interpretation of the previous result. In the Green and Porter theory, *unanticipated* demand shocks cause price wars. In the Rotemberg and Saloner theory, price wars are more likely when demand is high and anticipated to decline (see the next section). That the unanticipated component of demand, V_{1t} , is significant while the full demand shock U_{1t} is not allows us to distinguish between these explanations in favor of that of Green and Porter where Porter (1985) could not.

It is important to remember that the causes of price wars have been estimated from very few examples. The timing of these price wars is also not clear, with the “standard” model above not matching the onset date of Ulen’s (1979) classification for any of the price wars. For this reason, it is advisable to worry about whether the results above are very sensitive to the timing of the regimes. Table 4 presents the results of simple logit

TABLE 4 Price War Causes Using Reported Regimes

Regimes	$\text{Prob}\{J_{t+1} = 1 \mid J_t = 1, Z_t\} = \frac{e^{\gamma w_t}}{(1 + e^{\gamma w_t})}$			
	Model			
Parameter	1	2	3	4
<i>CONSTANT</i>	3.68 (.65)	3.55 (.65)	3.25 (.61)	2.92 (.57)
<i>BIGSHARE1</i>	-.71 (.46)			
<i>BIGSHARE2</i>		-.51 (.39)		
<i>BIGSHAREQ</i>			-.33 (.44)	
<i>SMALLSHARE</i>				.05 (.37)

* Note: Estimated standard errors in parentheses.

models that attempt to predict the probability of the onset of a price war in Ulen's classification, J_t . The results are surprisingly similar to those of the previous table, providing further evidence of the existence of a Green and Porter-type trigger and perhaps suggesting that with a better classification the results would be stronger.

Although the results described above support the applicability of the Green and Porter theory, I would like to point out that they are not completely positive. In a Green and Porter cartel it is not sufficient for a signal of cheating to increase the likelihood of a price war. It must be the case that the relationship is strong enough to deter firms from cheating. Absent data on costs and firm-specific elasticities, we can give only a very rough calculation of whether the proper incentives are maintained. Suppose that firm i offers a secret price cut in period t and raises its demand by 10%. If the price cut that induces this response is small, profits in period t may be increased by, say, 8–10%. How large is the strategic disincentive to cheat given the estimated effect of *BIGSHARE1*? The measure of demand, $\max_k(\log q_{kt} - 1/n \sum_j \log q_{jt})/\sigma_k$, would be expected to be increased by

$\log(1.1)/\sigma_i$ whenever firm i would otherwise have had the highest market share deviation. A simple computer simulation with reasonable parameter values gives the probability of a price war increasing by about .001. If the losses from a price war are about 15 times the per-period profits, the future loss is about 1–2% of current profits. The effect we have estimated is then several times too small to support collusion, and cannot be much larger given the precision of our estimates.

There are a number of possible explanations for this result. First, it is quite possible that we have significantly underestimated the responsiveness of the data to signals of cheating. This could happen because the firms have information not contained in our data, because imperfect price data induce misclassifications of the onset of price wars, and/or because a true trigger for price wars is only imperfectly correlated with *BIGSHARE1*. On the other hand, it could also be simply that the responses of the JEC were not sufficiently strong to deter cheating in a noncooperative equilibrium. While the results of this section support the applicability of the Green and Porter theory, it cannot be said that collusive equilibrium strategies have been clearly identified.

3. The Rotemberg and Saloner theory

■ **Theory and previous results.** The Rotemberg and Saloner model of collusion with time-varying demand is often misinterpreted simply as implying that price wars are likely

to break out during booms. While this is understandable given the title of the article, the Rotemberg and Saloner model has no uncertainty about past actions, and therefore has no price wars in the sense of the dramatic regime shifts of the Green and Porter model. Instead the theory is really one of countercyclical pricing, or more precisely, countercyclical degrees of collusion in the sense that firms set prices further from the monopoly level when demand is high than when it is low. The intuition for this is simple. Suppose firms are not very patient so that they are just able to deter cheating at a specified price. If demand in the current period is then increased relative to demand in the future, the gain from cheating will be increased relative to the cost of an ensuing punishment. In order to eliminate an incentive to cheat, the price must be lowered.

To apply rigorously the ideas of Rotemberg and Saloner's model to the JEC, it would be necessary to add time-varying demand to a Green and Porter-style model with unobservable prices. The optimal equilibrium of a model with imperfect observation is in one sense a very natural place to apply Rotemberg and Saloner's ideas because the incentive constraints do bind. Hence firms' incentives to cheat must be reduced whenever demand is expected to decline. However, in adapting a cartel design to a situation of declining demand, a variety of changes are possible. Prices in each phase may be raised or lowered, price wars may be lengthened or shortened, and the sensitivity of triggers may be adjusted. Prices will be countercyclical if the optimal adjustment involves lowering prices when demand is low, and price wars will be more common during booms if the optimal adjustment makes triggers more sensitive (or if the cartel fails to sufficiently reduce the incentive to cheat). In the absence of a formal analysis, it is possible to appeal to the literature by regarding booms as analogous to a lowering of firms' discount factors. The numerical solutions reported in Porter (1983a) then suggest that we should expect both countercyclical prices and more price wars during booms, although the results may be sensitive to such assumptions as the distribution of the error terms. In fairness to the Rotemberg and Saloner theory, it should be pointed out that the analysis is probably more applicable to situations less complicated than the one described above. While the effects do appear in the optimal equilibrium regardless of complexity, we could easily imagine that a nineteenth-century railroad cartel would instead choose a design with prices somewhat below the maximum possible, but in which prices and punishment thresholds need not be fine-tuned to the vagaries of each week's demand.

To apply the theory it is also necessary to reexamine what a "boom" means. A boom in the theory is a period when high current demand will be followed by a period of lower demand during an ensuing price war (Haltiwanger and Harrington, 1991). In the case of the JEC, we can identify two situations in which current demand is likely to be significantly higher than the future demand (with "future" here meaning the next fifteen or so weeks). First, given the serial correlation of demand shocks that has been estimated, a large random component of demand yesterday leads to increased demand today with less than a proportional increase in the future. This effect is closely related to the standard view of booms, and it justifies the conclusion without the assumption that demand shocks are observable in advance. Second, we have the entirely new interpretation that there may be a seasonal pattern to collusion. In the winter, when the Great Lakes were closed, the JEC received much higher demand. The incentive to deviate is thus at its peak just before the lakes are to melt. Given the parameter estimates of our standard model, each of the two effects should be of approximately the same strength. To summarize, an analysis of the applicability of the Rotemberg and Saloner theory should assess whether prices are less collusive or price wars are more likely to occur during booms, where booms are redefined as noted above.

Only two previous articles, those of Rotemberg and Saloner (1986) and Hajivassiliou (1989), have considered the Rotemberg and Saloner theory in the context of the JEC, and neither directly addresses any of the predictions outlined above. Rotemberg and Saloner

attempt only a rough look at annual frequencies of price wars and note that price wars are more common in years in which grain production is high and the Great Lakes remain frozen for a long time. These variables, however, do not identify situations in which demand is higher than it is expected to be in the next few months. Using only annual data, price wars also cannot be tied to the conditions of the week in which a price war begins. Hajivassiliou similarly considers only an index of grain production, and his analysis concerns not whether this index tends to cause price wars, but rather whether it tends to be high during price wars. No previous work has explored countercyclical pricing in the JEC.

□ **Model.** Conceptually, modifying the model of Section 2 to test the predictions of the Rotemberg and Saloner model is fairly simple. A variable B_t that measures the ratio of current to future demand should be added in in two places. To test whether B_t makes price wars more likely we can simply add B_t to the logit model for the transition probability $\text{Prob}\{I_{t+1} = 1 \mid I_t = 1, Z_t\}$. To test whether prices are less collusive when B_t is high we can use the modified supply equation

$$\log P_t = \beta_0 + \beta_1 \log Q_t + \beta_2 I_t + \beta_3 B_t I_t + \beta_{4-7} DMx_t + U_{2t}.$$

A negative coefficient on β_3 would indicate prices that are less collusive during booms.

In trying to perform such a test, the only difficulty lies in the determination of an appropriate variable B_t . To combine the effects of serially correlated shocks and seasonal factors I take B_t not to be any observable, but rather a function of the predetermined variables and parameters. The primary function used for this purpose is

$$BOOM_t = K \times \frac{\exp(\alpha_2 LAKES_t + \alpha_{3-14} SEASXX_t + \rho U_{1t-1})}{\sum_{s=1}^{52} (\hat{p}_{w,t+s} - \hat{p}_{c,t+s}) \exp(\alpha_2 LAKES_{t+s} + \alpha_{3-14} SEASXX_{t+s} + \rho^{s+1} U_{1t-1})} - 1.$$

The expression represents a ratio of present demand to the expected future demand during the course of a subsequent price war. Here, the parameters α_2 through α_{14} and ρ are those being simultaneously estimated while the value $\hat{p}_{w,t+s} - \hat{p}_{c,t+s}$ is the additional probability of being in a price war s periods after one begins implied by the results of the standard model.³

To further analyze the effect of booms, I will also separate the effects of serially correlated errors and seasonal variation. To analyze the first, I simply include $U_{1t-1} I_t$ in the supply equation and U_{1t} in the transition probability. Among the possible effects tested, the occurrence of price wars when U_{1t} is large is most similar to the common interpretation of the Rotemberg and Saloner model as implying price wars during booms. To look at seasonal variation, I create a function similar to that used for the combined effect,

$$SEASONAL_t = K \times \frac{\exp(\alpha_2 LAKES_t + \alpha_{3-14} SEASXX_t)}{\sum_{s=1}^{52} (\hat{p}_{w,t+s} - \hat{p}_{c,t+s}) \exp(\alpha_2 LAKES_{t+s} + \alpha_{3-14} SEASXX_{t+s})} - 1.$$

□ **Results.** While we might have expected that finding countercyclical levels of collusion would be easy given that it is not hampered by the small number of price wars, the results on the Rotemberg and Saloner model are somewhat disappointing, and we can make few firm conclusions given the imprecision of the estimates. The first set of columns in Table 5 gives the results of the estimation of a model in which the combined measure $BOOM_t$ affects the level of collusion and the likelihood of a price war. The estimate of

³ These latter values are not estimated simultaneously, as doing so would require a fixed point calculation to overcome the circularity of $BOOM_t$, depending on $\hat{p}_{w,t+s}$, which is in turn a function of $BOOM_t$. It is hoped that this will not unduly bias the estimates and standard errors.

TABLE 5 Effect of Booms

Variable	Combined Effect		Demand Shocks		Seasonal	
	Estimate	Standard Error	Estimate	Standard Error	Estimate	Standard Error
Demand						
CONSTANT	7.319	3.060	7.964	1.345	8.016	1.970
log P	-1.996	2.055	-1.594	.898	-1.574	1.459
LAKES	-.041	.127	-.036	.133	.008	.110
SEAS1	-.149	.181	-.045	.104	-.065	.125
SEAS2	.151	.155	.164	.139	1.136	.191
SEAS3	.236	.170	.080	.193	.124	.298
SEAS4	.100	.239	-.065	.185	-.048	.254
SEAS5	-.206	.279	-.312	.173	-.318	.181
SEAS6	-.451	.287	-.445	.207	-.513	.392
SEAS7	-.335	.283	-.298	.264	-.385	.504
SEAS8	-.371	.328	-.314	.330	-.440	.577
SEAS9	-.227	.426	-.265	.238	-.338	.416
SEAS10	-.340	.445	-.444	.185	-.531	.243
SEAS11	-.370	.297	-.478	.158	-.524	.203
SEAS12	-.311	.195	-.162	.162	-.403	.174
Supply						
CONSTANT	-6.856	6.801	-6.186	3.415	-4.419	2.545
log Q	.517	.673	.450	.335	.274	.250
DM1	-.223	.148	-.205	.090	-.164	.087
DM2	-.303	.179	-.302	.106	-.257	.070
DM3	-.437	.314	-.402	.148	-.326	.103
DM4	-.238	.130	-.319	.104	-.234	.197
I_t	.758	.416	.712	.195	.622	.119
$BOOM_t I_t$	-.233	.428				
$U_{t-1} I_t$			-.156	.164		
$SEAS_t I_t$.041	.417
Regimes						
CONST. ($I_t = 1$)	3.596	.474	4.604	.775	3.925	1.241
$BOOM_t (I_t = 1)$.446	1.582				
$U_{t-1} (I_t = 1)$			-.777	.477		
$SEAS_t (I_t = 1)$					-2.786	7.704
CONST. ($I_t = 0$)	2.624	.404	-2.620	.390	-2.677	.546
Other						
σ_1	.301	.061	.280	.041	.278	.064
σ_{12}	-.018	.023	-.017	.013	-.007	.004
σ_2	.204	.045	.187	.079	.155	.046
U_{t-1}	.838	.085	.831	.078	.824	.063
Log-likelihood	186.2		187.1		181.2	

the coefficient of $BOOM_t I_t$ in the supply equation indicates a countercyclical degree of collusion in pricing, with an implied decrease of about 3% in the collusive price when the incentive to deviate is one standard deviation above the mean. However, the standard error is almost twice this large. As a result, we cannot reject that collusion is procyclical or that the reduction is 12%. It appears that this boom effect is hard to separate econometrically from the effect of quantity on price. The second and third sets of columns examine the decomposition into seasonal and demand shock components in hopes that the models will be more precisely estimated or that we might gain confidence from a coincidence of estimates. The estimate of the countercyclical response to demand shocks is

TABLE 6 Causes of Price Wars

Parameter	Model		
	1	2	3
<i>CONSTANT</i>	3.64 (.49)	4.35 (.90)	3.87 (.71)
<i>BOOM_t</i>	.36 (1.48)		
<i>U_{1t}</i>		1.17 (1.15)	
<i>SEASONAL_t</i>			-2.51 (3.50)
<i>V_{1t}</i>		-5.00 (3.15)	

Note: Estimated standard errors in parentheses.

more precise, though still not significant, indicating a 5% drop in prices when U_{1t-1} is one standard deviation above the mean. No evidence at all is found for countercyclical price responses to seasonal cycles.

In hopes of obtaining better estimates of the effect of booms on the likelihood of price wars, Table 6 presents estimates from a model like that described above but with the indicators for booms entering only into the transition probabilities. As with the tests of the Green and Porter theory, there is little change in the structural parameters, so only the logit coefficients for the start of a price war are reported. The first model in the table includes our composite measure of booms and provides no support for the hypothesis that price wars are more likely during booms. Again, I decompose the boom effect into demand shock and seasonal components, because it is certainly possible that a cartel would choose to lower prices in some situations and make triggers more sensitive in others. The second model in the table is the repetition of a model from Table 3 and shows that the predictable pattern of decreasing demand after a positive demand shock does not cause price wars once we control for Green and Porter effects. Price wars do not appear to be more likely during booms as booms have previously been interpreted. The estimate of the seasonal effect is negative but fails to be significant.

The experience of the JEC is not obviously supportive of the hypothesis that price wars occur during booms as Rotemberg and Saloner had hoped after their initial analysis. However, given the lack of power in the tests, we cannot rule out the possibility that booms lead to a small degree of countercyclical pricing or a small increase in the frequency of price wars.

4. Secret price cuts

■ At this point, I return to an implication common to the Green and Porter and Rotemberg and Saloner theories, namely that firms do not offer secret price cuts. There are three main motivations for this effort. First, the fact that firms do adhere to the announced prices is an important prediction of each of the theories discussed. Second, while the mixed results of Section 2 may simply reflect the failure to identify the proper trigger, they might also indicate that no sufficiently strong trigger exists. In that case, profit-maximizing firms would be expected to have taken advantage of this and hence we might be able to detect fairly frequent cheating. Finally, the previous analysis has been based on list prices, which are presumed accurate because firms should not offer secret price cuts in equilibrium. If cheating took place, we might want to reconsider earlier tests.

There are no standard theoretical predictions about the nature of secret price cuts other than that they should not occur. Secret price cuts are also by nature not recorded in the JEC data. Given this combination, I should warn the reader that this section will be quite speculative. I would also like to mention that given reports in the contemporary trade press, it is very likely that some price cutting occurred.⁴ The real goal of this section is to see whether we can obtain some idea of the frequency and magnitude of such behavior so that we may comment on its implications. With these caveats, I move on to discuss methodology.

Suppose a secret price cut were offered only once for one week. An econometrician, like the JEC members, would have no chance of finding it. On the other hand, if periods of both price cutting and strict adherence to the official price are common, there is hope. In this case, it is as if secret price cutting is a significant omitted variable in the demand equation, shifting demand upward relative to that expected given the official price. Motivated by this observation, I follow a two-step approach. First, I examine the demand equation in the JEC for evidence of omitted variables. Subsequently, when the presence of such variables is indicated, I ask whether they may indeed reflect secret price cuts.

□ **Looking for secret price cuts via hidden regimes.** The problem of identifying secret price cuts is in many ways similar to that of identifying collusive and noncollusive behavior. In Section 2, collusion was identified by functional form assumptions—the residuals in the supply equation were assumed to be normal while collusive price increases are assumed in accordance with the Green and Porter theory to have a distribution with a two-point support. Here, such an approach is probably less appropriate given the lack of a theory describing the distribution of price cuts. If price cuts are normally distributed, they cannot be identified. It is probably true, though, that price cuts are not normally distributed, but instead have a mass point on no cut being offered with some continuous distribution over small positive price cuts. The approach taken here is simply to look for the easiest thing to find, a distribution of price cuts with a small finite support. Such price cuts would appear in the data as occasional shifts in demand relative to the official price. As the whole exercise is doomed if there is no evidence of any omitted variables in demand, I will look for such shifts first and worry later about whether they are secret price cuts.

To formally explore the possibility that omitted variables shift demand, I apply a model of hidden regimes with Markov transitions that is more complicated and somewhat more far-fetched than that of Section 2. Essentially, I add indicator variables for one or two additional demand regimes to the demand relation (1). The model I refer to as model 2 allows for two demand regimes so that the demand equation is

$$\log Q_t = \alpha_0 + \alpha_1 \log P_t + \alpha_2 LAKES_t + \alpha_{3-14} SEASXX_t + \alpha_{15} REGIME1_t + U_{1t}. \quad (5)$$

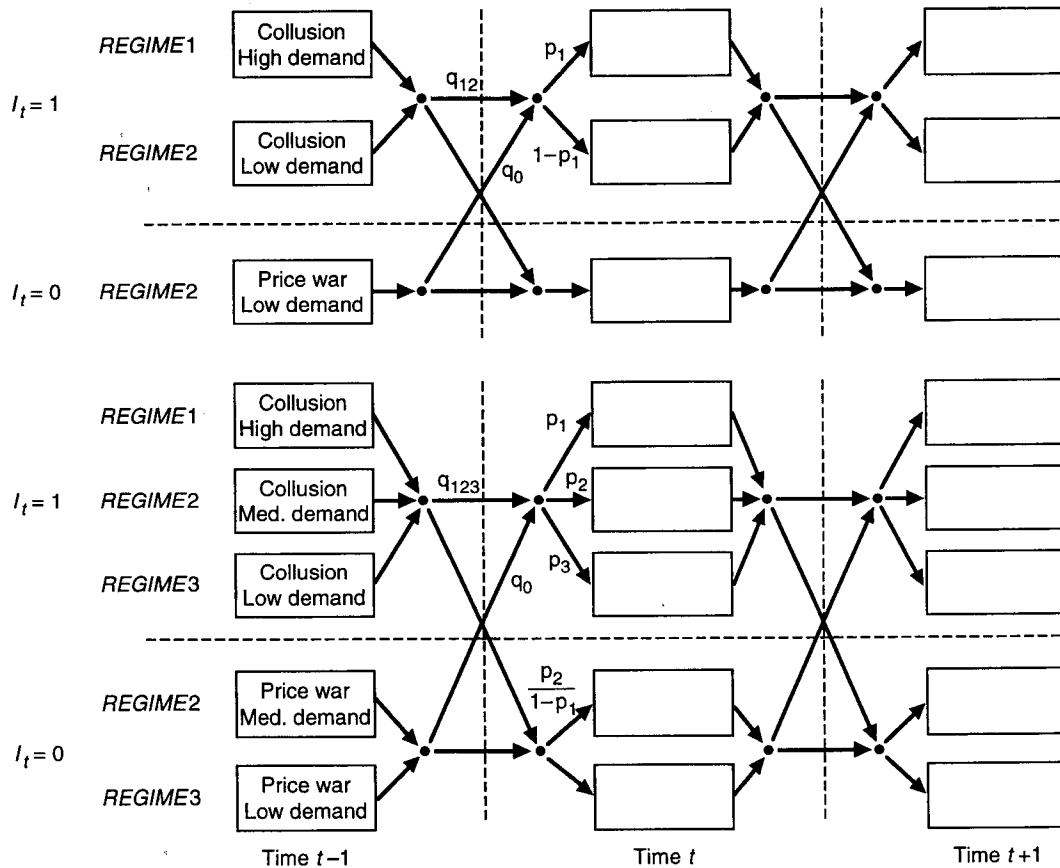
I intend for *REGIME1* to be an indicator for periods of price cuts, so I assume that this high-demand state occurs with probability p_1 whenever firms are trying to collude, but never occurs during price wars when firms are already presumably pricing at marginal cost.

Model 3 allows for three demand regimes so that the demand equation becomes

$$\begin{aligned} \log Q_t = & \alpha_0 + \alpha_1 \log P_t + \alpha_2 LAKES_t + \alpha_{3-14} SEASXX_t \\ & + \alpha_{15} REGIME1_t + \alpha_{16} REGIME2_t + U_{1t}. \end{aligned} \quad (6)$$

⁴ The *Daily Commercial Bulletin* of June 16, 1881, contains a typical report: "Some agents claim to insist on 20c Grain to New York, while others state that 15c has no doubt been accepted." An article from February 2, 1886, expresses more confidence: "There is little doubt but rates are being cut on shipments of Grain to the East by one or two lines."

FIGURE 1



Again I am envisioning that *REGIME1* will be an indicator for a high-demand regime due to secret price cuts. The lowest-demand regime might represent a period of unusually low demand due to severe weather, labor problems, unusually low steamship rates, etc. As it is not crucial to the problem at hand, I will not try to provide an interpretation for this regime. In a collusive state, I assume that the high demand *REGIME1* arises with probability p_1 , the medium-demand *REGIME2* with probability p_2 , and the third regime with probability $1 - (p_1 + p_2)$. I again assume that *REGIME1* never arises during a price war and that the other two regimes occur with the same relative probabilities under collusive and price war pricing.

Figure 1 illustrates the structure of regime transitions in models 2 and 3. The large boxes represent the different regimes possible at time t . In model 2, there are two regimes with collusive prices (the two possible demand regimes) and one with competitive prices. The arrows indicate probabilistic transitions and are labelled with the probabilities with which they are taken. For example, the high-demand *REGIME1* occurs with probability p_1 , and the low-demand *REGIME2* occurs with probability $1 - p_1$ whenever the firms are colluding. Within each of the regimes, prices and quantities are determined by the supply and demand relations (3) and (5), with the appropriate values of the regime dummies, e.g., with $REGIME1_t = 1$ and $I_t = 1$ within the uppermost period- t state in the figure. The diagram for model 3 is similar, but with the five states representing the three possible demand regimes under collusive pricing and the two possible demand regimes in a price war.

The estimation of these models is almost unreasonably successful. Maximum-likelihood estimates are reported in Table 7. Most structural parameter estimates are quite similar to those of our standard model. Many of the key parameters, including the price

TABLE 7 Estimates of Hidden Regimes

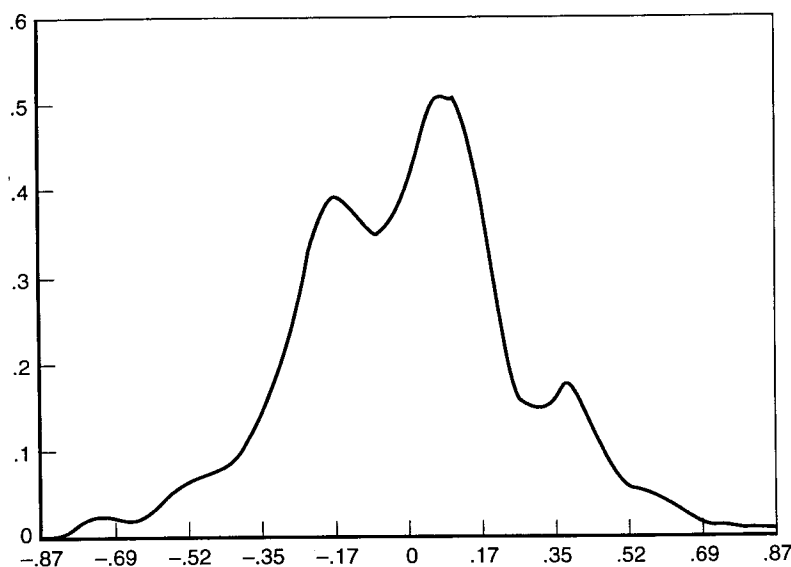
Demand: $\log Q_t = \alpha_0 + \alpha_1 \log P_t + \alpha_2 LAKES_t + \alpha_{3-14} SEASXX_t + \alpha_{15} REGIME1_t + \alpha_{16} REGIME2_t + U_{1t}$
 Price: $\log P_t = \beta_0 + \beta_1 \log Q_t + \beta_2 I_t + \beta_{3-6} DMX_t + U_{2t}$

Variable	Model 2		Model 3	
	Estimate	Standard Error	Estimate	Standard Error
Demand				
CONSTANT	8.147	.870	7.651	.176
$\log P$	-1.354	.711	-1.587	.092
LAKES	-.222	.073	-.229	.092
SEAS1	-.046	.122	-.114	.087
SEAS2	.166	.147	.163	.095
SEAS3	.016	.115	.106	.108
SEAS4	.056	.137	-.041	.122
SEAS5	-.225	.169	-.321	.127
SEAS6	-.271	.158	-.374	.162
SEAS7	-.163	.194	-.139	.159
SEAS8	-.350	.205	-.332	.157
SEAS9	-.179	.182	-.184	.167
SEAS10	-.215	.122	-.219	.148
SEAS11	-.194	.121	-.089	.128
SEAS12	-.216	.083	-.091	.114
REGIME1	.439	.036	.728	.027
REGIME2			.384	.022
Supply				
CONSTANT	-3.861	1.577	-4.101	.412
$\log Q$.220	.153	.242	.038
DM1	-.165	.069	-.162	.072
DM2	-.244	.062	-.239	.068
DM3	-.331	.069	-.323	.063
DM4	-.228	.082	-.198	.119
I_t	.595	.079	.608	.047
Transition probabilities				
q_{123}	.979	.012	.976	.037
q_0	.061	.028	.070	.076
p_1	.650	.071	.256	.083
p_2			.519	.086
Other				
σ_1	.121	.015	.089	.006
σ_{12}	-.001	.001	-.0002	.0007
σ_2	.097	.021	.099	.011
U_{1t-1}	.888	.031	.947	.018
Log-likelihood	529.0		582.5	

elasticity of demand and the effect of collusion on prices, are more significant, particularly in model 3. There is also a truly dramatic increase in fit relative to the standard model, with an improvement in log-likelihood similar to that reported by Porter when he first added a second supply regime to the model. The most striking result is the extremely high degree of significance of the dummy variables *REGIME1* for the hidden demand regimes we were looking for. The *t*-statistics on the *REGIME1* parameter estimates, 12.1 and 26.7 in models 2 and 3 respectively, are almost unreasonably high given a sample size of 328 weeks. In model 2, the high-demand *REGIME1* has demand increased by about 55%. In model 3, the high-demand *REGIME1* has demand 41% above the level of the medium-demand *REGIME2*, while the low-demand regime will have demand 32% below that of *REGIME2*.

At this point, I think it is important to caution the reader that it is not claimed that we have estimated the effect of secret price cuts. The high-demand regimes of the two

FIGURE 2



models are very different, with that of model 2 occurring in 65% of the periods and that of model 3 occurring in 26% of them. Clearly, at least one model is finding evidence of something other than price cutting. I conclude only that there is very strong evidence of significant omitted variables in the demand equation.

To help understand these results it is instructive to look at the demand residuals of the standard model. If no price cuts were given, these might appear to be normally distributed. If occasional price cuts of the same magnitude were given, the distribution would have an additional peak shifted to the right by an amount equal to the demand response to a price cut of the given magnitude. Figure 2 presents a kernel density estimate of the residuals \hat{V}_{1t} from the standard model. I think the reader will agree that the residuals do not appear to be normally distributed. The smaller peak to the right resembles the predicted effect of price cutting and is the empirical regularity that accounts for the high-demand *REGIME1* in model 3.

□ **Do the estimated regimes look like secret price cuts?** The fact that I have so far ignored the problem of interpreting the results of the previous model should not be taken to imply that I do not see it as important. To the contrary, I wish to emphasize strongly that the reported results could easily be attributed to a number of other factors (e.g., weather, irregular schedules, steamship competition, mechanical problems) or simply to departures from the functional form or distributional assumptions. I believe that the interpretation of models with hidden regimes is an important and often overlooked question and therefore I discuss it at some length in this section.

Secret price cuts are just one of a number of possible explanations for our findings, and we cannot claim to know much about many of the alternate explanations. We cannot then simply perform hypothesis tests of competing theories to decide among them. Instead, I will try to provide a number of small analyses to see whether the omitted variables have the properties we would expect of secret price cuts.

The first type of evidence I consider is that which has been most common in other studies—a simple look at the reasonableness of the parameter estimates. For example, Porter (1983b) compares the magnitude of his estimated supply shift to the prediction that “collusive” supply shocks would increase prices by something less than, but of the same magnitude as, the difference between monopoly and competitive prices. We have little

theory to go on. Perhaps the best standard for comparison we have is obtained from reports in the contemporary trade press which indicate that perhaps 10–25% is a reasonable size for price cuts.⁵ As far as the frequency of price cuts goes, we can only say that the fact that the cartel went as long as two years between price wars suggests that price cuts were not too common.

In model 2, the high-demand *REGIME1* occurs with estimated probability $\hat{p}_1 = .65$ when firms are colluding. Given the estimated price elasticity of 1.35, a regime shift of the indicated magnitude would only result from price cuts of between 25% and 30%. It seems unlikely that price cuts of this size could have occurred so often without clear historical evidence and more frequent price wars. I think that this alone is sufficient evidence to conclude that the high-demand regime of model 2 does not reflect secret price cutting, and will examine the regimes in this model no further. In model 3, the high-demand *REGIME1* occurs with estimated probability $\hat{p}_1 = .26$. Given the estimated price elasticity of 1.59, the increase in demand from *REGIME2* to *REGIME1* would require a price cut of about 20%. This certainly seems like a reasonable size and frequency for price cuts, although given our lack of precise predictions about price cutting and alternative explanations, we cannot be much more confident that we have found the effect of price cuts.

The second type of evidence I would like to consider is the relationship between the demand shifts I have found and the available historical evidence on price cutting. The argument can be thought of as analogous to Porter's use of the similarity between his estimates and Ulen's reported regimes to justify the price war interpretation. In the case of secret price cuts, two relationships to the historical data might be envisioned. First, if secret price cuts cannot be kept completely secret, we might expect the high-demand regime to be accompanied by simultaneous reports in the press of price cutting. Such reports are likely to be highly inaccurate, however, because shippers who receive special treatment will want to keep it secret to maintain their advantage in the grain market, and those who do not will gladly claim they did if it might trigger a price war. Second, to the extent that the JEC members believe such reports, whether true or not, we might expect also that secret price cutting will break out in the periods immediately following the reports.

Historical evidence on price cutting was compiled from the Chicago Board of Trade's *Daily Commercial Bulletin*, which contains both price quotes for rail freight transportation and occasional reports of rumored price cuts. Two indicator variables were created. The dummy variable *CUTREPORTED* is set to one in any week in which an article stated that secret price cuts were rumored to be given. The dummy variable *CUTLISTED* is set to one whenever the average of the daily prices listed in the newspaper was at least three cents below the official price. The two variables identify twelve and eighteen weeks respectively as involving price cuts.

Table 8 presents a simple descriptive analysis of the relationship between the maximum-likelihood classification of weeks into demand regimes and three potential indicators of collusion: the two noted above and Ulen's series of reported price wars. Each row gives the frequency with which the appearance of the indicator at time t is accompanied by the occurrence of the high-demand regime in periods t through $t + 5$. The fractions are computed over the sample of periods for which periods t and $t + s$ both involve collusive pricing. Entries marked with an asterisk indicate that the high-demand regime is significantly more common when the indicator occurs than in the rest of the sample.⁶ (Remember, about 26% of all periods involve the high-demand regime.) The first two columns of

⁵ Some such reports appear in the *Daily Commercial Bulletin* of June 16, 1881, September 25, 1884, and January 22, 1885.

⁶ The standard errors have been computed conditional on the regime classification, not taking the estimation of the regimes into account.

TABLE 8 Historical Evidence on Secret Price Cuts

	Indicator of Price Cuts (C_t)		
	<i>CUTREPORTED</i>	<i>CUTLISTED</i>	Reported Price War
$E(\text{REGIME1}_t C_t, I_t)$.44	.30	.25
$E(\text{REGIME1}_{t+1} C_t, I_{t+1})$.57*	.13	.28
$E(\text{REGIME1}_{t+2} C_t, I_{t+2})$.66*	.33	.40*
$E(\text{REGIME1}_{t+3} C_t, I_{t+3})$.50	.40	.46*
$E(\text{REGIME1}_{t+4} C_t, I_{t+4})$.33	.60*	.41*
$E(\text{REGIME1}_{t+5} C_t, I_{t+5})$.20	.25	.34

* The high-demand regime is significantly more common when the indicator occurs than in the rest of the sample.

the table are based on very small samples, so only very large differences will be significant.

In the first line of the table, we see that the indicators are slightly more likely to be accompanied by the appearance of the high-demand regime, though not significantly so. In the next several lines we see that the high-demand regime is much more likely to occur in the few weeks following the occurrence of reports of cheating, in accordance with the speculation that firms might engage in retaliatory price cutting. The results of this section would likely be even stronger if it were not for the fact that one-third to one-half of the time a price war has begun within a month of the reports and hence the points are dropped from the frequency calculation in the lower rows of the table. Approximately 15% of the occurrences of the high-demand regime are directly attributable to excess occurrences of the high-demand regime following one of the price cutting indicators. All together, the table provides fairly convincing evidence that the high-demand regime estimates reflect, at least in part, secret price cutting.

If the JEC is in fact a Green and Porter-style cartel in which firms react to aggregate demand shocks, there is also a fairly direct test available of whether the hidden regimes are or are not observable to the JEC members. Absent Rotember and Saloner-type effects, we would expect that the trigger for a Green and Porter cartel would be based only on demand shocks unobservable to the firms. If the demand regimes are perfectly observable, we would expect a trigger to be based on the residuals of the model, i.e., on V_{1t} . If the demand regimes are completely unobservable, an aggregate demand trigger would be based on the sum of the demand residual and the effect of the regime shift variables.

FIGURE 3

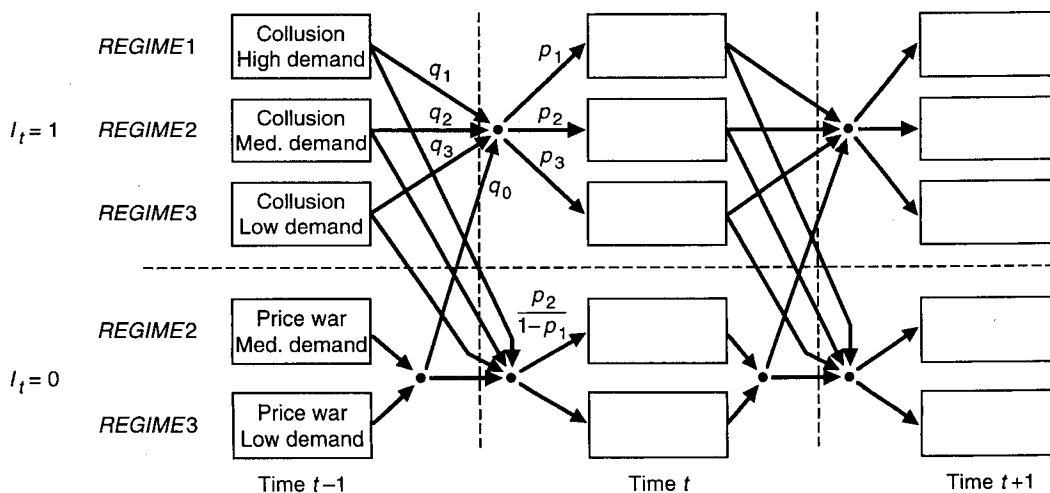


TABLE 9 Observability of Hidden Regimes

Demand: $\log Q_t = \alpha_0 + \alpha_1 \log P_t + \alpha_2 LAKES_t + \alpha_{3-14} SEASXX_t + \alpha_{15} REGIME1_t + \alpha_{16} REGIME2_t + U_{1t}$
 Price: $\log P_t = \beta_0 + \beta_1 \log Q_t + \beta_2 I_t + \beta_{3-6} DMX_t + U_{2t}$
 Regimes: $q_i = \frac{e^{\gamma W_t}}{(1 + e^{\gamma W_t})}$ $W_t = (1 V_{1t} R_t)$

Variable	Estimate	Standard Error
Demand		
CONSTANT	7.67	.48
log P	-1.57	.43
LAKES	-.23	.06
SEAS1	-.11	.04
SEAS2	.16	.12
SEAS3	.10	.13
SEAS4	-.04	.10
SEAS5	-.32	.09
SEAS6	-.38	.09
SEAS7	-.14	.10
SEAS8	-.33	.10
SEAS9	-.19	.11
SEAS10	-.22	.07
SEAS11	-.09	.07
SEAS12	-.09	.06
REGIME1	.73	.04
REGIME2	.38	.05
Supply		
CONSTANT	-4.07	.97
log Q	.24	.09
DM1	-.16	.07
DM2	-.24	.06
DM3	-.32	.07
DM4	-.20	.07
I _t	.61	.07
Transition probabilities		
q _i CONST.	5.65	1.06
q _i V _{1t}	-3.17	1.26
q _i R _t	-.48	1.03
q ₀	.07	.03
p ₁	.26	.05
p ₂	.52	.07
Other		
σ ₁	.09	.01
σ ₁₂	.0001	.001
σ ₂	.10	.02
U _{1t-1}	.947	.02
Log-likelihood	584.4	

To test between these hypotheses, I estimate a model like model 3 but in which the transition from a collusive to a price war state is a function both of V_{1t} and of $R_t \equiv \alpha_{15} REGIME1_t + \alpha_{16}(REGIME2_t - 1)$. Figure 3 illustrates the nature of the regime transitions in the model. The probabilities q_1 , q_2 , and q_3 of continuing in a price war are all assumed to be determined by a single logit model with V_{1t} and R_t as explanatory variables. (Note that each of these quantities is dependent on the demand regime.) With completely unobserved regimes, V_{1t} and R_t should have identical coefficients. With observed regimes, the coefficient on R_t should be zero. While it may seem completely unreasonable to suppose that we could identify the form of the trigger given that we could barely say that a trigger exists, the results in Table 9 do in fact yield significant results. At the 5%

level we can reject the equality of the coefficients on V_{1t} and R_t in the logit model of transition probabilities. We therefore reject the idea that the regimes represent only factors unobservable to the JEC members. To the contrary, the estimates suggest that the high-demand regime is mostly due to factors the JEC recognized and which did not lead them to initiate price wars.

In this section we have seen that there are significant unobserved variables affecting demand, and that one of these regimes present in one-quarter of all periods is compatible with price cuts of about 20%. While there is historical evidence of price cutting correlated with the estimated regimes, it is likely that secret price cutting accounts for only a fairly small fraction of these periods. These findings are obviously not compatible with the prediction of the theory that no cheating should take place, but also they are not indicative of the JEC's having provided grossly inadequate incentives to cooperate. The results of this section can also be seen as strengthening the conclusions of Section 2. Viewing the model in Table 9 as one that partly corrects for the presence of omitted variables in demand, we have found that aggregate demand shocks not only still appear to cause price wars, but that the results are now clearly statistically significant.

5. Conclusion

■ Previous work testing the Green and Porter theory has focused on the prediction that price wars should occur. Using the familiar data on the Joint Executive Committee, this is verified again here, with the traditional view challenged only in that I find prices to be much closer to the monopoly level than has been previously reported. More importantly, I have used a model that focuses on the regime transition probabilities to provide evidence for the existence of the trigger strategies that are central to the theory. Given the serially correlated demand shocks, such trigger effects can be distinguished from the predictions of the Rotemberg and Saloner model. The fact that several triggers are found to be significant and significant results are also obtained using an alternate classification is suggestive of the existence of a stronger mechanism than has been identified.

The results are not completely positive, in that the initial estimates are not highly significant and do not identify the presence of a mechanism sufficiently strong to deter cheating. The literal prediction of the theory that no secret price cutting should occur also appears to be false. The case for the existence of a stronger trigger is perhaps strengthened, though, by the fact that cheating is not too widespread.

In the context of cartels with imperfect observations, the Rotemberg and Saloner theory has a number of new implications. In the case of the JEC, booms relevant to this theory must be understood to be short-term demand shocks or seasonal peaks in demand. Booms may lead to less collusive prices and/or more frequent price wars. I am unable to find either of these effects, and in particular do not find evidence for the common interpretation of price wars being more likely during booms. My estimates are disappointing in their lack of power, however, so moderate effects are certainly possible. In any case, the JEC is also perhaps among the less likely places one would want to apply the Rotemberg and Saloner theory, so these results may not be reflective of the theory's applicability in other contexts.

References

- ABREU, D., PEARCE, D., AND STACCHETTI, E. "Optimal Cartel Equilibria with Imperfect Monitoring." *Journal of Economic Theory*, Vol. 39 (1986), pp. 251–269.
- BERRY, S. AND BRIGGS, H. "A Non-parametric Test of a First-Order Markov Process for Regimes in a Non-cooperatively Collusive Industry." *Economic Letters*, Vol. 27 (1988), pp. 73–77.
- BRESNAHAN, T.F. "Empirical Studies of Industries with Market Power." In R. Schmalensee and R.D. Willig, eds., *The Handbook of Industrial Organization*. New York: North-Holland, 1989.

- COSSLETT, S.R. AND LEE, L.F. "Serial Correlation in Latent Variable Models." *Journal of Econometrics*, Vol. 27 (1985), pp. 79-97.
- Daily Commercial Bulletin*. Chicago: Howard White & Co., 1880-1886.
- FUDENBERG, D., LEVINE, D., AND MASKIN, E. "The Folk Theorem with Imperfect Public Information." Mimeo, Harvard University, 1993.
- GREEN, E.J. AND PORTER, R.H. "Noncooperative Collusion Under Imperfect Price Competition." *Econometrica*, Vol. 52 (1984), pp. 87-100.
- HALTIWANGER, J., AND HARRINGTON, J.E. "The Impact of Cyclical Demand Movements on Collusive Behavior." *RAND Journal of Economics*, Vol. 22 (1991), pp. 89-106.
- HAIJIVASSILIOU, V. "Testing Game-theoretic Models of Price-fixing Behavior." Cowles Foundation Discussion Paper no. 935, 1989.
- LEE, L.F. AND PORTER, R.H. "Switching Regression Models with Imperfect Sample Separation Information—With an Application on Cartel Stability." *Econometrica*, Vol. 52 (1984), pp. 391-418.
- MACAVOY, P.W. *The Economic Effects of Regulation*. Cambridge, Mass.: MIT Press, 1965.
- PORTER, R.H. "Optimal Cartel Trigger Price Strategies." *Journal of Economic Theory*, Vol. 29 (1983a), pp. 313-338.
- . "A Study of Cartel Stability: The Joint Executive Committee, 1880-1886." *Bell Journal of Economics*, Vol. 14 (1983b), pp. 301-314.
- . "On the Incidence and Duration of Price Wars." *Journal of Industrial Economics*, Vol. 33 (1985), pp. 415-426.
- ROTEMBERG, J.J. AND SALONER, G. "A Supergame-Theoretic Model of Price Wars During Booms." *American Economic Review*, Vol. 76 (1986), pp. 390-407.
- STIGLER, G.J. "A Theory of Oligopoly." *Journal of Political Economy*, Vol. 72 (1964), pp. 44-61.
- ULEN, T.S. *Cartels and Regulation: Late Nineteenth Century Railroad Collusion and the Creation of the Interstate Commerce Commission*. Ph.D. dissertation, Stanford University, 1979.